An Empirical Investigation of Structural Breaks in the Ex Ante Fisher Effect*

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Abstract

This paper investigates the relationship between expected inflation and the nominal interest rate using Australia data. Recently developed time series techniques are used that allow for estimation across different regimes where the timing and number of structural breaks are not known *a priori*. The results are consistent with the existence of significant structural breaks in the relation between interest rates and inflation, with there being some evidence that these are associated with changes in taxation. After allowing for the structural breaks, it appears that interest rates fail to fully reflect anticipated inflation.

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1 Introduction

The hypothesis that the nominal rate of interest reflects anticipated changes to the price level is usually attributed to Irving Fisher (1930).[1] Fisher's hypothesis states that the one-period nominal interest rate on a financial asset is the sum of two components. The first is the ex ante real interest rate; it represents a borrower's expected return from acquiring a financial asset once allowance is made for the second component; the anticipated change in a general index of goods prices over the course of the period.

A case can be made for the existence of a "one-for-one" relation between the nominal rate of interest and the expected rate of inflation. This requires the real interest rate to be determined by the deep structural parameters that condition technology and tastes, and a belief that these deep structural parameters change only slowly. It is in this form that the Fisher hypothesis is usually tested, essentially examining whether or not the real rate of interest is orthogonal to the expected rate of inflation. The most common finding in the empirical literature is a rejection of the hypothesis, with the data often showing the nominal interest rate failing to adjust fully to price changes (see Crowder and Hoffman 1996 for a survey of the relevant research).

A variety of empirical techniques have been used to test the Fisher hypothesis. Most recently, cointegration and error correction methodologies have been used to investigate whether (i) the nominal interest rate and some measure of expected inflation cointegrate with a unitary coefficient (evidence that the

[1] Antecedents for this view can be found in Masrshall (1938) and in Thorton's *Nature of the Paper Credit* (Blaug 1978).

strong form Fisher hypothesis describes an empirically valid long-run equilibrium relation), or (ii) the real interest rate is not cointegrated with the expected inflation rate (a necessary condition for the validity of the Fisher hypothesis). Recently, doubts have been raised whether these time series techniques are capable of robustness in the face of major structural changes in the economy (Gregory and Hansen 1996, Zivot and Andrews 1992). Possible candidates for structural changes of the magnitude that might affect empirical testing of the Fisher hypothesis are events such as a change in the monetary authority's operating procedures, or the implementation of major structural reforms such as financial deregulation. Events such as these might be expected to lead to parameter instability in the relation between the nominal interest rate and the expected rate of inflation, and thus call into doubt the findings of cointegration analyses based on a stable long-run equilibrium relation.

One response to the possibility of structural instability is to modify the cointegration procedure to make allowance for structural change, see Malliaropulos (2000). The difficulty here is that techniques which enable this, such as Zivot and Andrews' (1992) sequential unit root test, choose one structural break point only from within the sample. Whether such a technique would enable the identification of a common stochastic trend between expected inflation and the nominal interest rate if there is more than one structural break is debatable.

In this paper, the approach taken to testing the empirical validity of the Fisher effect differs from the previous research in three respects. First, the testing procedure looks explicitly at the *ex ante* Fisher effect. This is achieved by constructing a model-consistent set of expected inflation observations through the exploitation of the information content of nominal interest rates. Second, a consistent estimate of the parameter relating the nominal interest rate to expected inflation is derived from the structural restrictions placed on the bivariate error term in a VAR featuring the one-period ahead inflation rate and the nominal interest rate. Thirdly, a recently developed methodology devised by Bai and Perron (1998) is used which enables a test of Fisher's hypothesis to be carried out that is robust to multiple structural breaks at unknown points in the sample.

These techniques are applied to quarterly Australian data. Australia is a particularly interesting case since (i) different studies using essentially the same data have reached different conclusions about the Fisher hypothesis (Atkins 1989, Inder and Silvapulle 1993) and (ii) Australia had very stringent interest rate controls throughout the 1970s, a decade in which inflation peaked well into the teens, only to have these controls rescinded in the early 1980s. Instability in the relation between the nominal interest rate and expected inflation would therefore be unsurprising. In a previous paper, I showed that the strong form of the Fisher effect was rejected overall for Australia but that there was some evidence in support of the hypothesis if only the post-deregulation data were used (Olekalns 1996). However, in that paper, the timing of the break-point was imposed *a priori* on the data. As is well known, pre-testing in this way leads to a degradation of power when testing the null of no structural change. Hansen (2001) makes this point in a comprehensive survey of the relevant literature on the econometrics of structural change. However, in this paper, it is the data itself that will identify likely points of structural change in the sample.

The paper is organised as follows. The methodology and the data are described, respectively in sections 2 and 3. The results are presented and discussed in section 4. This is followed, in section 5, by some concluding remarks.

2 Method

Under the assumption of rational expectations, the Fisher effect can be described by the following equation;

$$R_t = \rho + \beta E(\pi_{t+1}|I_t) + \varepsilon_t, \tag{1}$$

where R_t is the nominal interest rate, ρ is a constant term, $E(\pi_{t+1}|I_t)$ is the expected inflation rate conditional on the information set available in the current period, and ε_t is an error term. Under the strong form of the Fisher hypothesis, the coefficient β will equal unity, meaning that the nominal interest rate fully incorporates anticipated changes to the price level and that ρ can be interpreted as the constant real rate of interest.

Two preconditions must hold for the strong form of the Fisher hypothesis to be valid. The first is that interest receipts must not be subject to taxation. Where this is not true, the Fisher equation should be written as,

$$(1 - \tau_t)R_t = \rho + \beta E(\pi_{t+1}|I_t) + \varepsilon_t, \tag{2}$$

where τ_t is the tax rate. Note that estimating equation (1) when equation (2) is correct will result in an overstatement of expected inflation's effect on the nominal interest rate. For the purposes of exposition, I will maintain the assumption that the tax rate is equal to zero in this section of the paper. A discussion of the empirical effects of introducing a non-zero tax rate into the analysis will be postponed until later.

The second precondition is the requirement that the interest elasticity of the demand for money be equal to zero (Sargent 1972). Sargent derived this result in the context of a dynamic Keynesian income expenditure model. The instantaneous effect of an increase in expected inflation is to stimulate the demand for investment as the real interest rate would have fallen. If the interest elasticity of the demand for money is non-zero, the subsequent rise in the nominal interest rate will, at least initially, be less than the increase in anticipated inflation. A zero interest elasticity, however, ensures that the nominal interest rate fully adjusts to the change in anticipated inflation.

Any test of equation (1) requires a measure of expected inflation. A common practice is to use the actual rate of inflation as a proxy. However, as long as β is non-zero, the nominal interest rate must have a component that reflects the expected inflation rate. As a result, the information content of the nominal interest rate can be exploited to derive a series for expected inflation. A vector autoregression having the following bivariate structure provides a means of doing this;

$$\upsilon_t = \sum_{i=1}^{q} A_i \upsilon_{t-i} + e_t, \tag{3}$$

where υ_t is the 2 x 1 vector of variables (π_{t+1}, R_t), and where e_t is a 2 x 1 vector of residuals. Each coefficient matrix, A_i, has dimensions 2 x 2.

It is convenient to write the vector autoregression in first-order form;

$$\begin{bmatrix} v_t \\ v_{t-1} \\ \cdot \\ \cdot \\ \cdot \\ v_{t-q+1} \end{bmatrix} = \begin{bmatrix} A_1 & A_2 & \cdot & \cdot & A_q \\ I_2 & 0_2 & \cdot & \cdot & 0_2 \\ 0_2 & I_2 & 0_2 & \cdot & 0_2 \\ \cdot & \cdot & & & \cdot \\ \cdot & & \cdot & & \cdot \\ 0_2 & \cdot & \cdot & I_2 & 0_2 \end{bmatrix} \begin{bmatrix} v_{t-1} \\ \cdot \\ \cdot \\ \cdot \\ v_{t-q} \end{bmatrix} + \begin{bmatrix} I_2 \\ 0_2 \\ \cdot \\ \cdot \\ 0_2 \end{bmatrix} e_t,$$

where I_2 is the 2 x 2 identity matrix and 0_2 is a 2 x 2 matrix of zeros. This first order system can be written more compactly as

$$Q_t = AQ_{t-1} + We_t, \tag{4}$$

where $Q_t = (v_t \; v_{t-1} \ldots v_{t-q+1})'$ and $W = (I_2 \; 0_2 \; 0_2 \; \ldots \; 0_2)'$. Based on this first order system, the expected value of inflation can be calculated from

$$E(\pi_{t+1}|I_t) = z'AQ_t \tag{5}$$

where z' is of length 2q and is defined by $z' = (1\;0\;0\;\ldots\;0)$. In effect, z' selects the forecast of inflation from $E(Q_{t+1}|I_t) = AQ_t$.

Given equations (1) and (5), the Fisher hypothesis can be written as

$$R_t = \rho + \beta z' AQ_t + \varepsilon_t, \tag{6}$$

where expected inflation is now explicitly measured using the VAR.

The information content contained in the innovations to, respectively, expected inflation and the interest rate can be exploited to derive a consistent estimate of the structural parameter, β (see Keating 1990). To see this, consider the expected interest rate in period t where the expectation is conditional on the information available in period $t-1$;

$$E(R_t|I_{t-1}) = \rho + \beta E(E(\pi_{t+1}|I_t)|I_{t-1}). \tag{7}$$

From the law of iterated expectations, equation (7) can be written as

$$E(R_t|I_{t-1}) = \rho + \beta E(\pi_{t+1}|I_{t-1}) = \rho + \beta z' A({}_{t-1}Q_t), \tag{8}$$

where ${}_{t-1}Q_t$ is the projected value of Q_t based on the information in period $t-1$. Therefore, innovations to the nominal interest rate can be written as

$$R_t - E(R_t|I_{t-1}) = \beta[z' A Q_t - z' A({}_{t-1}Q_t)]. \tag{9}$$

Noting that ${}_{t-1}Q_t = AQ_{t-1}$ and using equations (4) and (9), it follows that

$$R_t - E(R_t|I_{t-1}) = \beta z' A W e_t. \tag{10}$$

Equation (10) implies the existence of a structural relationship between the respective innovations to expected inflation and to the nominal interest rate. In effect, the arrival of new information is transmitted, via the Fisher effect, to the nominal interest rate, and this transmission is conditioned by the value of the parameters in A.

To derive an equation that is suitable for estimation, let ξ_t be a variable that incorporates innovations to the nominal interest rate in period t that are unrelated to innovations in expected inflation. These factors can be incorporated into the model by writing equation (10) as

$$R_t - E(R_t|I_{t-1}) = \beta z' A W e_t + \xi_t. \tag{11}$$

Finally, define a variable $S_t \equiv z' A W e_t$. If we interpret ξ_t as a regression error, a consistent estimate of β, can be found from the regression

$$R_t - E(R_t|I_{t-1}) = \beta S_t + \xi_t. \tag{12}$$

The variables on both the left and right sides of equation (12) are, by construction, stationary.

3 The Data

The 90 day Commercial Bill Rate is used as the measure of the nominal interest rate. The inflation rate is given by 400*ln($\mathrm{CPI}_t/\mathrm{CPI}_{t-1}$), where CPI_t is the value of the consumer price index in period t. The data are quarterly and span the period 1969:3 to 2000:1. The source of the data is the DX database.[2] The data are displayed in Figure 1.

- FIGURE 1 HERE -

4 Results

The first stage in the empirical analysis is to estimate the VAR, equation (3). The optimal lag length, using the Bayesian Information Criterion (BIC) to guide selection of the lag length, is two quarters. [3] Figure 2 shows the actual values

[2] The respective DX mnemonics are GCPIAGU (for the CPI) and FIRMMBAB90 (for the commercial bill rate).

[3] The values of the BIC were respectively 9.578 (1 quarter), 9.540 (two quarters) and 9.614 (three quarters).

for inflation and the fitted values from the inflation component of the VAR. The fit is reasonably close; the adjusted R^2 is 0.48.

- FIGURE 2 HERE -

The results from OLS estimation of equation (12) are shown in Table 1.

- TABLE 1 HERE -

The Wald test, which is based on a heteroskedastic consistent estimate of the standard error, shows that the strong form of the Fishe hypothesis, $\beta = 1$, cannot be rejected over the full sample. However, a series of diagnostic tests reveal that these results are almost certainly affected by specification error; the indications are that the error term is not normally distributed, suffers from serial correlation and is characterised by fourth order ARCH.

This specification error might be related to instability in the Fisher effect over the sample period. Although no such instability is revealed by a CUSUM test (see Figure 3), the results from a CUSUM squared test are indicative of instability (see Figure 4). A further indication that instability might be a problem is provided by Figure 5, which shows sequential estimates of the parameter β. The first of these estimates is derived by truncating the sample at the twentieth observation, estimating (3) over the first twenty observations, then deriving the necessary series to calculate the right and left hand sides of (12). An OLS regression is then run to obtain an estimate of β. This procedure is then repeated, each time adding an additional observation until the full sample estimate is derived. The results of this sequential estimation procedure suggest that there may have been a major structural change in the value of the β parameter around the early 1980s.

- FIGURE 3 HERE -

- FIGURE 4 HERE -

- FIGURE 5 HERE -

In light of this *prima facie* evidence of structural instability, a recent technique devised by Bai and Perron (1998) can be implemented which enables estimates to be made of multiple break points. The technique involves estimating m single equations allowing for, respectively, l, $l + 1$, . . . $l + m$ possible structural breaks. The estimated sum of squared residuals are then compared across the regressions and the global minimum value is established. If this value is sufficiently small relative to the estimated sum of squared residuals with fewer structural breaks, then that specification becomes the preferred model.[4]

The Bai and Perron procedure was implemented allowing for a minimum segment length of three years (twelve observations). The estimation procedure allows for first order autocorrelation of the residuals and for the variance of the residuals to differ across the segments.[5]

The results of applying Bai and Perron's procedure are shown in Table 2. The results in the table are consistent with there being two structural breaks, in

[4] Bai and Perron (1998) provide asymptoticly valid critical values for inferring the number of structural breaks.

[5] The GAUSS program to implement the program was supplied by Pierre Perron.

1975:3 and in 1980:2.[6] Table 3 shows the OLS estimates of β for the respective sub-periods. The standard errors are calculated using a heteroskedastic and autocorrelation consistent covariance matrix. These estimates are consistent with the strong version of the Fisher effect holding in the first sub-period (i.e., the null hypothesis, $H_0 : \beta = 1$, cannot be rejected at the five percent level of significance). However, in the middle sub-period, 1975:3-1980:1, the relation between the nominal interest rate and the expected rate of inflation breaks down completely with β being insignificantly different from zero. The relation between expected inflation and the interest rate is restored in the third sub-period; however, β, although significantly positive, is less than one (based on a one-tailed test of significance).

- TABLE 2 HERE -

- TABLE 3 HERE -

A variety of causal factors could be responsible for this instability. One of these factors, often overlooked, is the effect that changes in the tax rate levied on interest income may have on the Fisher effect. As outlined previously, in an economy in which interest receipts are taxed, the coefficient connecting the nominal interest rate with the expected rate of inflation is $\beta/(1-\tau_t)$ where τ_t, the tax rate, can vary over time. Estimates of β, using pre-tax interest rates, will exhibit instability if derived from a sample period during which the tax rate changed.

[6] These dates are consistent with the more informal indications of structural breaks provided by the CUSUM squared test and by the recursive estimation of β.

Since the majority of bonds in Australia are held by corporations (Inder and Silvapulle 1993), it is the corporate tax rate that is most relevant in this context. There have been significant changes in the Australian corporate tax rate over the sample period. Figure 6 shows two measures of the corporate tax rate. The first is the marginal corporate tax rate which has changed ten times over the sample period. The second is calculated as the ratio of corporate tax paid to capital income and is best thought of as the average tax rate applicable to corporations. This displays much more variation than the marginal tax rate, being affected by changes in the denominator (i.e., income) as well as by changes in allowable deductions and other strategies that corporations might use to influence their tax liability.[7]

- FIGURE 6 HERE -

To examine whether tax changes are responsible for the instability in the Fisher effect, the procedures described above were applied to equation (2). That is, a bivariate VAR was estimated, but this time featuring the after-tax rate of interest. An estimate of β can then be derived from equation (12) that will be invariant to changes in the tax rate. It could reasonably be inferred that the instability identified previously in the Fisher effect does not have its origins in changes to the tax system if estimates based on equation (12) continued to reveal parameter instability when using the after-tax interest rate. Conversely, should equation (12) be stable over the sample period using the post-tax interest

[7] The data for these tax rates come from the Dx Database. The marginal tax rate is series VNEQ.UN_RTC and the average tax rate is series VTEQ.AR_RTK.

rate, then tax changes are likely to have been the source of the instability found with the pre-tax interest rate.

Table 4 shows the full sample estimates derived from equation (12) using two definitions of the post-tax interest rate, the first using the marginal tax rate and the second using the average tax rate. Under neither tax rate is the strong form of the Fisher hypothesis supported; Wald tests of the null hypothesis $H_0 : \beta = 1$ show that the null is easily rejected. However, the diagnostics suggest that the results derived using the average tax rate are to be preferred. There is evidence of non-normality in the residuals as well as serial correlation and ARCH effects for the marginal tax rate. For the results derived using the average tax rate, the only diagnostic test that the specification fails is the test for the normality of the residuals. Examination of the residuals suggests that there may be an outlier at the eighteenth observation. Repeating the estimation with a dummy variable for that observation removes the non-normality (the JB Normality test statistic is 1.222 with a p-value of 0.543) and yields an estimate of β equal to 0.449 (with a standard error of 0.090). A Wald test shows that this is significantly different from unity.

- TABLE 4 HERE -

The results from Bai and Perron's procedure, using the average tax rate to calculate the post-tax interest rate, are shown in Table 5. The first two test statistics, the UDMax and WDMax statistics, suggest that there may have been at least one structural break, although, at the five percent level of significance, the values of the test statistics are very close to the respective critical values.

The various SUPF tests are consistent with there being only one break, which occurs in the second quarter of 1975. However, on the basis of the BIC, we would conclude that there had been no breaks during the sample period. Therefore, the evidence in favour of structural breaks is a good deal more tenuous when the post-tax interest rate is used compared to the results obtained when changes to the tax rate are ignored. This provides some evidence that the instability in the Fisher relationship is driven by tax changes, and not, as is commonly argued (Olekalns 1996) by the effects of financial deregulation.[8]

- TABLE 5 HERE -

5 Conclusion

This paper has tested the validity of an equation based on the Fisher effect as an accurate representation of the relation between the nominal interest rate and inflation for Australia. Previous studies had found that the decision whether to reject the Fisher equation was sensitive to the sample period used (and in particular, whether pre- or post-deregulation data were used). The results in this study, however, reject the strong form of the Fisher effect even when structural breaks in the relation between the nominal interest rate and inflation are allowed for in the estimation procedure. Unlike other studies, however, allowance for structural breaks was made using a technique which did not require an *a priori*

[8] The possibility that the instability might be due to instability in the VAR equations themselves was investigated by running Bai and Perrons' procedure on the VAR equations. The results revealed no instability for the inflation equation. However, there was a possible structural break in the equation for the interest rate (at observation #19). Estimating β allowing the coefficients on the interest rate VAR equation to change at observation #19 yielded an estimate of 0.6696 with a heteroskedastic consistent standard error of 0.238. Comparing this to the results in Table 1 show that allowing for the potential instability in the VAR makes no significant difference to the results.

decision to be made about the timing of breaks. Evidence was also presented which was suggestive that the structural breaks identified were related to the taxation of interest receipts.

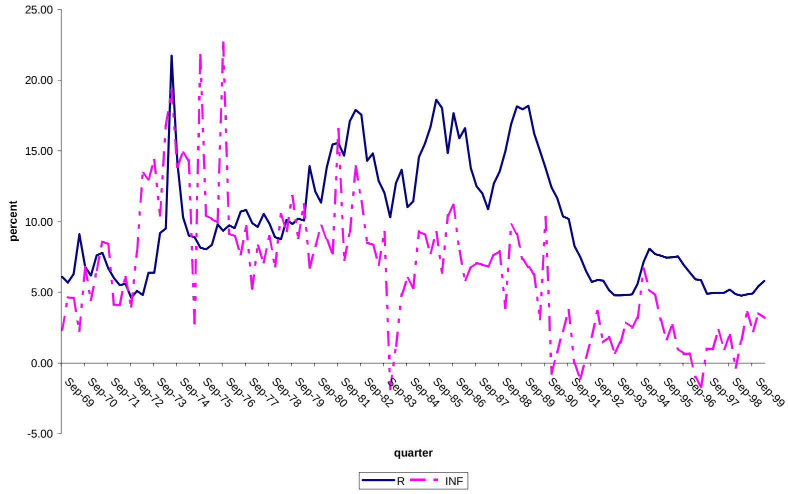


Figure 1: The Data

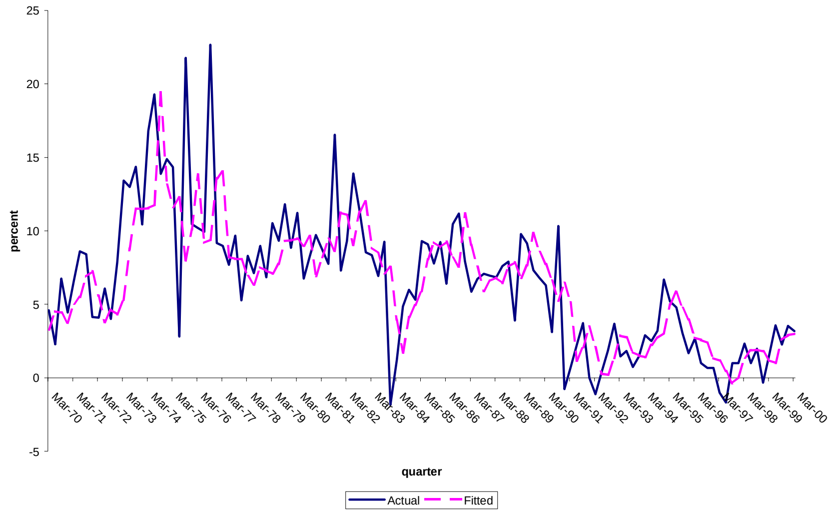


Figure 2: Actual and Fitted Inflation Data

Table 1: Full Sample Estimates

[p-values]

β	0.698
SE	0.235
Wald ($\beta = 1$)	1.659 [0.198]
JB Normality Test	84.231 [0.000]
Breusch-Godfrey Test	10.890 [0.028]
ARCH LM Test	15.176 [0.004]

$Note$: Newey-West standard errors are heteroskedastic consistent and are calculated using the technique of Newey and West (1987). The Wald test is of the null hypothesis that $\beta = 1$. The JB normality test is the Jarque-Berra (1980) test. The Breusch-Godfrey Test is an LM test of the null of no serial correlation (up to the fourth order); see Godfrey (1988). ARCH LM test for ARCH of up to fourth order; see Engle (1982).

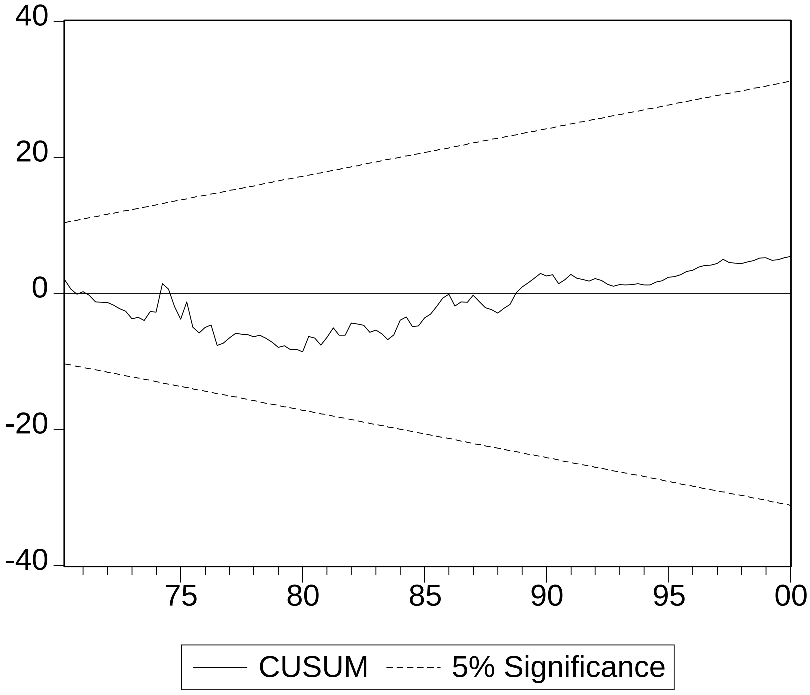
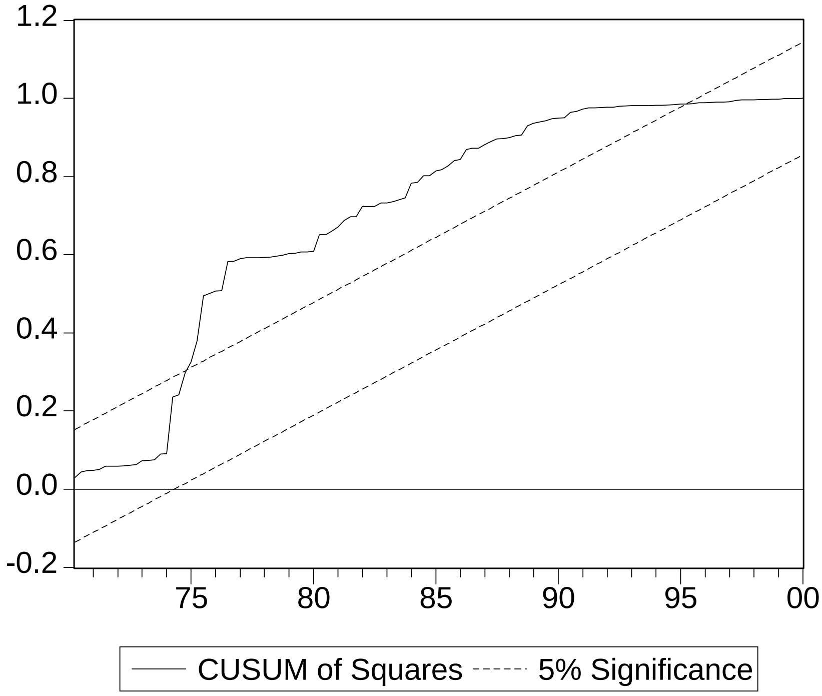


Figure 3:



1.2
1.0
0.8
0.6
0.4
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-0.2
75
80
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90
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00
CUSUM of Squares
5% Significance

Figure 4:

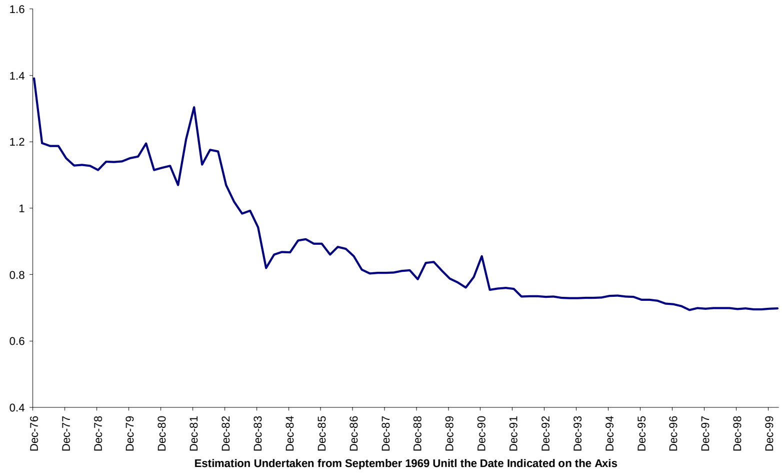

Figure 5: Sequential Estimates of β

Table 2: Multiple Structural Breaks

UDMax	19.342
WDMax	29.190
SupF(2|1)	26.724
SupF(3|2)	5.944
SupF(4|3)	5.609
SupF(5|4)	3.338
# of Breaks (BIC)	2
Break Dates	1975:3
	1980:2

Notes: UDMax and WDMax are tests of the null hypothesis of no structural change against the alternative of some unknown number of break points. The 5% critical values are, respectively, 9.520 and 10.390. SupF(i+1|i) is a test of the null hypothesis of i structural changes against the alternative of i+1 structural changes. The 5% critical values for i = 1, 2, ... , 5 are respectively, 9.100, 10.550, 11.360 and 12.350. # of Breaks (BIC) identifies the number of breaks that minmises the value of the Schwarz information criterion.

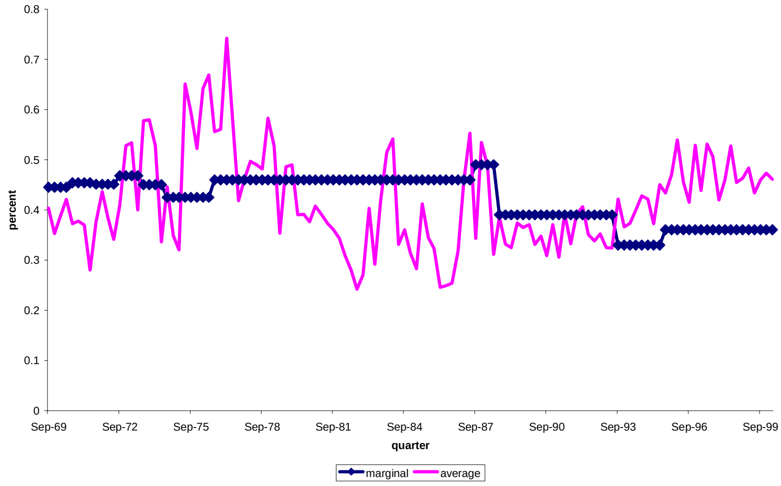

Figure 6: Corporate Tax Rates

Table 3: Estimates of β

Period	β	Standard Error
1970:1-1975:2	1.272	0.319
1975:3-1980:1	0.076	0.046
1980:2-2000:1	0.687	0.108

Table 4: Estimates of β

(post-tax interest rate)

[p-values]

	marginal tax rate	average tax rate
β	0.366	0.569
SE	0.131	0.162
Wald ($\beta = 1$)	23.560 [0.000]	7.094 [0.009]
JB Normality Test	78.662 [0.000]	105.443 [0.000]
Breusch-Godfrey Test	9.749 [0.045]	3.665 [0.530]
ARCH LM Test	11.613 [0.020]	2.001 [0.736]

Note: See notes to Table 1.

Table 5: Multiple Structural Breaks

(post-tax interest rate)

UDMax	9.778
WDMax	13.485
SupF(2|1)	9.577
SupF(3|2)	9.176
SupF(4|3)	5.649
SupF(5|4)	4.226
# of Breaks (BIC)	0
Break Dates	1975:2

Notes: See notes to Table 2.

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